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Document de Travail
Working Paper

EA 4272

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2011/04

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Testing the Speculative Efficiency Hypothesis on CO_2 Emission Allowance Prices: Evidence from Bluenext*

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*The authors would like to thank the Conseil Français de l'Energie for their financial support of this research.

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Abstract

In this paper, we attempt to examine the speculative efficiency hypothesis on CO_2 emission allowance prices negotiated on Bluenext, by testing the relationship between futures and spot prices from the Fama (1970) framework. This approach is based on the joint hypothesis of no risk premium and unbiasedness of futures prices. Cointegration tests are performed to confirm the legitimacy of futures and spot prices being included in the regression, following the approach proposed by Balke and Fomby (1997). The results indicate the absence of linear and nonlinear cointegration relationship between spot and futures prices. The speculative efficiency hypothesis did not hold even if the joint hypothesis is not rejected because of the existence of serial correlation in the residuals.

Keywords: CO_2 emission allowances; Cointegration; Spot and futures prices; Market efficiency.

JEL Classification: G13; G14; Q50.

1 Introduction

On January 2005, the European Union Emission Trading Scheme (EU ETS) went into effect. The EU ETS is one of the most important initiatives taken to reduce the greenhouse gas (GHG) emissions (primarily CO_2) that cause climate change (Kyoto protocol). The European Union (EU)[1] includes approximately 11,500 participating installations spread across twenty-seven member states. In 2010, it is estimated that the sources to which the trading scheme applies account for 45 per cent of CO_2 emissions and a little less than 40 per cent of total GHG emissions in that year.

The EU ETS introduces a cap-and-trade system, which operates through the creation and distribution of tradable rights to emit, usually called EU allowances (EUAs), to installations. Since a constraining cap creates a scarcity rent, these EUAs have value. The distribution of these rights for free is called allocation and is the unique feature of cap-and-trade system. The cap-and-trade scheme operates over discrete periods, with the first or pilot period (Phase I, 2005-2007) and with the second period corresponding to the first commitment period of the Kyoto Protocol. This period will extend from 2008 to 2012 (Phase II) and will be followed by a third period from 2013 to 2020 (Phase III). The EU target is a reduction of 8 per cent below 1990 emissions in the 2008-2012 period. To help countries in achieving their reduction objectives, the Protocol includes three flexibility mechanisms: The creation of an international carbon market, Joint Implementation and the Clean Development Mechanism.[2]

The EU ETS includes spot, futures, and option markets with a total market value of

[1] To improve the fluidity of the EU ETS, organized allowance trading has been segmented across trading platforms: Nordic Power Exchange (Nord Pool) in Norway began in February 2005, European Energy Exchange (EEX) in Germany began in March 2005, European Climate Exchange (ECX) based in London and Amsterdam started in April 2005, BlueNext in France and Energy Exchange Austria (EEA) in Austria began in June 2005, and SendeCO2 in Spain started at the end of 2005.

[2] Joint Implementation (JI) projects do not create credits, but rather transfer reduction units from one country to another. The aim of Clean Development Mechanism (CDM) projects is to promote investments in developing countries by industrialized nations and to encourage the transfer of low-emission technologies.

72 billion Euros in 2010. Futures contracts account for a wide part of this value (about 87% in 2010).. Understanding the relationship between spot and futures prices is thus of crucial importance for all participants in the carbon market. Carbon trading works only if markets for carbon provide enough liquidity and pricing accuracy, i.e., markets provide prices that are useful for hedgers and other users of carbon markets. The efficiency of the CO_2 market is particularly important for emission intensive firms, policy makers, risk managers and for investors in the emerging class of energy and carbon hedge funds.

As pointed out by Fama (1970), a financial market can be considered as efficient if prices fully reflect all available information and no profit opportunities are left unexploited known as the weak-form efficiency or speculative efficiency hypothesis. Speculative efficiency is essential for an operator who wants to hedge on the futures market against any possible price fluctuations. According to the futures markets literature, the model that futures prices are unbiased estimators of future spot prices is the appropriate framework to test efficiency. Using this model, efficiency will necessarily imply that the market price fully reflects available information and thus there exists no strategy that traders can speculate in the futures market on the future levels of the spot prices exploiting profits consistently. One way to test the link between spot and futures prices is to use cointegration tests.[3] In others words, it should expect spot and futures prices for any commodity to be linked through a long-run equilibrium relationship because it can be argued that spot and futures prices are driven by the same fundamentals. Evidence of no cointegration seems to be consistent with the speculative market hypothesis. Indeed, economic theory suggests that cointegration is unlikely to be observed in efficient markets (Granger, 1986). Another way to test whether futures prices are the best predictors of the spot prices is to use the regression proposed by Fama (1970). In this approach, market efficiency requires that futures prices should be unbiased predictors of future spot prices and

[3]Cointegration is a necessary condition for market efficiency but nor a sufficient condition. A test for serial correlation is needed to infer about market efficiency.

residuals of the regression are not serially correlated. Simple empirical tests are based on the joint hypothesis of market efficiency and unbiasedness of futures prices (risk neutrality).

Although relevant papers have been published on the behavior of emission allowance spot and futures prices (see, e.g., Alberola et al., 2008; Daskalakis and Markellos, 2008; Paolella and Taschini, 2008; Seifert et al., 2008; Benz and Trück, 2009), studies on CO_2 market efficiency are rather sparse. Alberola and Chevallier (2009), Uhrig-Homburg and Wagner (2009) and Joyeux and Milunovich (2010) study the relationship between spot and futures prices from the cost-of-carry model. However, Daskalakis and Markellos (2008) and Daskalakis et al. (2008) have some doubts on the applicability of a cost-of-carry model because of the EU ETS is a very young market means that significant differences in terms of stakeholders, liquidity, information and pricing may exist between spot and futures markets. Recently, Chevallier (2010) analyzes the relation between spot and futures prices using daily data from February 26, 2008 to April 15, 2009.[4] He rejects a cointegrating relationship between CO_2 allowances spot and futures prices when accounting for the presence of a structural break in February 2009, possibly due to the delayed impact of the "credit crunch" crisis. A vector autoregression analysis indicates that futures prices are relevant for price information in the spot market (the opposite is not true). Charles et al. (2011) examine the martingale difference hypothesis (MDH) within the EU ETS during the Phase I and the Phase II, using both daily and weekly data over the 2005-2009 period. Their results indicate that for the Phase I, the spot price changes are predictable, suggesting the possibility of abnormal returns through speculation. The authors fail to rejet the MDH during the Phase II, meaning that the European carbon market seems to be weak-form efficient.

The aim of this paper is to investigate the speculative efficiency hypothesis between spot and futures prices in order to provide empirical evidence for efficiency on the EU

[4] Note that Chevallier (2010) employ the daily spot prices negotiated on BlueNext and the daily futures prices negotiated on ECX when investigating the relation between spot and futures prices.

ETS. We thus study the daily spot et futures prices negotiated on Bluenext, covering the period from February 22, 2008 to October 20, 2010, namely the Phase II.[5] We test the joint hypothesis of market efficiency and unbiasedness of futures prices from the Fama (1970) regression. As it is well known that appropriate tests for efficiency and unbiasedness are necessarily dependent upon the underlying time-series properties of the data, we first use various unit roots tests, and we then apply linear (Johansen, 1995) and nonlinear (Seo, 2006) cointegration tests following the approach proposed by Balke and Fomby (1997).

The remainder of this paper is organized as follows: Section 2 presents the speculative efficiency hypothesis. The empirical framework is discussed in Section 3. The conclusion is drawn in Section 4.

2 The Speculative Efficiency Hypothesis

Theoretically, if spot and futures markets operate efficiently and are frictionless, futures contracts should be traded at a price known as the fair value. The starting point of most studies is the arbitrage free or cost-of-carry model in which the futures price is represented as

$$F_t = S_t e^{(r+u-d)(T-t)} \tag{1}$$

where F_t is the futures price at time t; S_t is the spot price at time t; r is the risk-free interest rate; u is the storage cost; d is the convenience yield; and T is the expiration date of the futures contract and $(T-t)$ is the time of expiry of the futures contract. Taking logarithms of both sides of equation (1) gives

$$\mathrm{Ln}(F_t) = \mathrm{Ln}(S_t) + (r+u-d)(T-t) \tag{2}$$

[5] BlueNext is the most liquid spot CO_2 exchanges operating under the EU ETS. Contrarily to others studies (Uhrig-Homburg and Wagner, 2009; Chevallier, 2010) we take the futures prices negotiated also on Bluenext but not on ECX, which is the most liquid futures exchange for EUAs, to have the data negotiated on the same market.

This equation suggests that the long-term relationship between the logs of the spot and futures prices should be one to one.

In practice, researchers have had difficulty testing the arbitrage relationship embodied in equation (2) due to the unobservable nature of storage costs and convenience yields (Joyeux and Milonovich, 2010). Hence, most studies have focused on the Fama (1970) speculative market efficiency tests of the form

$$S_t = \alpha + \beta F_{t-1} + \varepsilon_t \qquad (3)$$

In this approach, market efficiency requires that futures prices should be unbiased predictors of future spot prices. Simple empirical tests of the speculative efficiency hypothesis are based on tests of the joint hypothesis $\alpha = 0$ and $\beta = 1$. Failure to reject the joint hypothesis implies that the futures price determined at time $t-1$ is an unbiased predictor of the spot price for time t. However, statistical rejection of this joint hypothesis means either that there is a risk premium ($\alpha \neq 0$) or that the market is inefficient ($\beta \neq 1$). As underlined by Bilson (1981), it is important to note, however, that best unbiased forecasting by the futures price is not a necessary component of an efficient markets approach. It is easy, for example, to construct a model in which markets are efficient in the sense of removing any opportunities for risk-adjusted excess returns but in which there is a predictable bias in the futures price forecast.

It is well known that appropriate tests for efficiency and unbiasedness are necessarily dependent upon the underlying time-series properties of the data. If the price series are non-stationary, hypothesis tests based on equation (3) will give biased results. Regressing a non-stationary variable, which can only be made stationary by differencing on a deterministic trend, generally leads to the problem of a spurious regression, involving invalid inferences based on t- and F-tests (Granger and Newbold, 1974). In such cases the researcher could falsely conclude that a relationship exists between two unrelated non-stationary series. One way to circumvent the stationary problem is to estimate equation (3) in first-difference form (Hansen and Hodrick, 1980)

$$\Delta S_t = \gamma + \delta \Delta F_{t-1} + \varepsilon_t \qquad (4)$$

where Δ is the first-difference operator. Market efficiency and unbiasedness are jointly implied by the restrictions $\gamma = 0$ and $\delta = 1$.

Nevertheless, it is well known that equation (4) are misspecified if the two series (spot and futures prices) are cointegrated, that is to say if they have the same stochastic trend in common. When two price series, such as the future and the spot price series, are both integrated of the same order d, a linear combination of two $I(d)$ series can be integrated of an order lower than d. More specifically, it is possible that two series that are non-stationary and contain a unit root, for example $I(1)$, can generate a linear combination that is stationary, $I(0)$. These two series are said to be cointegrated with a cointegrating relationship of the following form

$$\varepsilon_t = S_t - \alpha - \beta F_{t-1} \tag{5}$$

If the two series are cointegrated, i.e. showing a stable common relationship in the long term, it is possible that the movement of one asset is linked to the movement of other asset. Thus, the establishment of a cointegration relationship is equivalent to the existence of an error correction term, which implies that in the face of a deviation of one asset price from the induced long-run relationship. Indeed this term describes the adjustment process due to disequilibrium. In the case of cointegration relationship between spot and futures prices, it is necessary to use an error correction representation described in Granger (1986)

$$\Delta S_t \quad = \quad -\rho\varepsilon_{t-1} + \theta\Delta F_{t-1} + \sum_{i=2}^{m} \theta_i \Delta F_{t-i} + \sum_{j=1}^{n} \psi_i \Delta S_{t-j} + \nu_t \tag{6}$$

where ε_{t-1} is the error correction term, and ν_t is a stationary white-noise residual term. Cointegration implies $\rho > 0$ because spot price changes respond to deviations from long-run equilibrium. The speculative hypothesis implies the following restrictions $\rho = 1$, $\theta \neq 0$ and $\theta_i = \psi_i = 0$.

3 Empirical results

The study sample consists of the daily closing prices of spot EUA prices from February 22, 2008 to October 20, 2010 (661 observations) and futures EUA prices of maturity

December 2010, December 2011 and December 2012 from April 21, 2008 to October 20, 2010 (624 observations) both negotiated on BlueNext.[6] Figure 1 provides a graphical representation of the spot and futures series. Note that the futures prices are higher than spot prices. This market condition is known as contango. Note that as the futures contract approaches to its maturity date, the difference between futures and spot prices is smaller and diminishes to zero at maturity since spot and futures prices converge.

We apply various unit root tests on spot and futures prices and find that all price series are characterized by a unit root (Table 1). When tests are applied on series in first-difference, they are found to be stationary. In other words, all price series are integrated of order 1. These results confirm those obtained by Alberola et al. (2008), Daskalakis et al. (2009), Chevallier (2009) and Alberola and Chevallier (2009).

Table 2 presents summary statistics for the returns calculated as the first differences in the logs of the EUA spot and futures prices. The kurtosis coefficient is significant for the both series, implying that the distribution of the log-returns is leptokurtic and thus the variance of the CO_2 prices is principally due to infrequent but extreme deviations. A leptokurtic distribution has a more acute peak around the mean and fat tails. The Lagrange Multiplier test for the presence of the ARCH effect clearly indicates that the log-returns show strong conditional heteroscedasticity, which is a common feature of financial data. In other words, there are quiet periods with small price changes and turbulent periods with large oscillations. Moreover, the skewness coefficient is negative and significant only for the spot series, implying that there is more negative log-returns than positive log-returns. This result means that the distribution of the spot price changes is asymmetric. No evidence of skewness is found in the futures log-returns.[7]

To test for cointegration between the spot and futures prices, both linear

[6] Data for Bluenext are available on www.bluenext.fr.

[7] Note that the spot prices display less volatility (measured by standard deviation) than do futures prices. This can suggest that the investors tend to be less conservative in their trading approach and take price shocks in the spot market seriously.

(Johansen,1995) and nonlinear (Seo, 2006) tests following the approach proposed by Balke and Fomby (1997) are used.

We first implement the Johansen maximum likelihood procedure (Johansen, 1988, 1991). This approach consists in estimating a Vector Error Correction Model (VECM) by maximum likelihood, under various assumptions about the trend or intercept parameters and the number r of cointegrating vectors, then conducting likelihood ratio tests. We write a p-dimensional VECM as follows

$$\Delta y_t = A' X_{t-1}(\beta) + u_t$$

with

$$X_{t-1} = \left\{ \begin{array}{ccccc} 1 & w_{t-1}(\beta) & \Delta x_{t-1} & \ldots & \Delta x_{t-n} \end{array} \right\}'$$

where x_t is a p-dimensional $I(1)$ time series which is cointegrated with one $(p \times 1)$ cointegrating vector β, $w_t(\beta) = \beta' x_t$ is the $I(0)$ error-correction term, u_t is an error term, and A is a coefficient matrix.

Johansen (1995) considers five restrictions on the deterministic components. In model 1 the level data y_t have no deterministic trends and the cointegrating equations do not have intercepts, giving the most restrictive specification. In model 2 the level data y_t have no deterministic trends and the cointegrating equations have intercepts. In model 3 the level data y_t have linear trends but the cointegrating equations have only intercepts. In model 4 the level data y_t and the cointegrating equations have linear trends. In model 5 the level data y_t have quadratic trends and the cointegrating equations have linear trends, giving the least restrictive specification. These five cases are nested from the most restrictive to the least restrictive. As noted by K*ü*hl (2007), the formulation of the model has important implications for testing the market efficiency hypothesis. To obtain the correct model, statistical inferences must be carefully examined first. A LR test is thus carried out (Johansen, 1994). The form of the LR tests is as follows

$$LR = -T \sum_{i=r+1}^{p} \ln \left[\frac{1-\hat{\lambda}_i^j}{1-\hat{\lambda}_i^k} \right] \qquad \text{with } j,k = 1,\ldots,5 \text{ and } j \neq k. \tag{7}$$

From the results reported in Table 3, we concluded that model 1 seems to be the most appropriated to test the cointegration relationship between the spot and futures price series.

Johansen (1988, 1991) proposes two LR test statistics to test whether there is no cointegration under the null against the alternative linear cointegration. The first test, called lambda max test, is based on the log-likelihood ratio

$$\mathrm{LR}(r|r+1) \quad = \quad -T\ln(1-\lambda_{r+1})$$

where T is the sample size, k is the number of endogenous variables, and $r = 0,1,\ldots,k-1$. This LR test tests the null hypothesis that the cointegration rank is equal to r against the alternative that the cointegration rank is equal to $r+1$.
The second test, called the trace test, is based on the log-likelihood ratio

$$\mathrm{LR}(r|k) \quad = \quad -T\sum_{i=r+1}^{k}\ln(1-\lambda_i)$$

where λ_i is the largest eigenvalue of the A matrix in equation, k is the number of endogenous variables, and $r = k-1,\ldots,1,0$. This LR test tests the null hypothesis of r cointegrating relations against the alternative of k cointegrating relations, where k is the number of endogenous variables, for $r = 0,1,\ldots,k-1$.
Results of these tests are given in Table 4.[8]

The null hypothesis of none cointegrating vector between the spot and futures prices cannot be rejected, implying that the linear VECM does not seem to be the most suitable model for the data of interest. This finding is in contradiction with that of Chevallier (2010). This difference can be explained by the fact that Chevallier (2010) employs (i) the daily spot prices negotiated on BlueNext and the daily futures prices negotiated on ECX when investigating the relation between spot and futures prices; and (ii) a shorter period (February 26, 2008 to April 15, 2009) than our period of interest. Evidence of no cointegration seems to be consistent with the speculative effi-

[8]We have to specify the lags of the test VAR to apply the Johansen cointegration tests. We use the traditional criteria (Akaike, Schwarz, Hannan-Quinn, Final Prediction Error) to select the optimal lag length.

ciency hypothesis.[9] Indeed, economic theory suggests that cointegration is unlikely to be observed in efficient markets (Granger, 1986). The Equation 6 means that the price of one asset does not only depend on its own past prices but also on the history of a different asset's prices, implying that the speculative efficiency hypothesis is violated (Richard, 1995).

Nevertheless, as pointed by Balke and Fomby (1997), the concept of cointegration is based on the implicit assumption that the adjustment of the deviations towards the long-run equilibrium is made instantaneously at each period. There are nevertheless serious arguments in economic theory to invalidate this assumption of linearity.[10] Moreover, in the linear cointegration context, increases or decreases of the deviations are deemed to be corrected in the same way. Again, several theoretical arguments may contest this assumption, such as the presence of menu costs (Levy et al., 1997), market power (Ward, 1982) or simply small country versus rest of the world effects. Balke and Fomby (1997) introduce the concept of threshold cointegration.[11] In their nonlinear framework, the adjustment does not need to occur instantaneously but only once the deviations exceed some critical threshold, allowing thus the presence of an inaction or no-arbitrage band. While their work focused on the long-run relationship representation, extension to a threshold VECM has been made by several authors, the threshold effect being applied to the anticipations by the agents of interventionary policy only to the error-correction term (Seo, 2006) or to the lags and the intercept as well (Hansen and Seo, 2002). Moreover, spurious cointegration can occur when there are breaks in the deterministic component (level or slope) of each time series

[9]Cointegration is a necessary condition for market efficiency but not a sufficient condition. A test for serial correlation is needed to infer about market efficiency.

[10]Among them, the presence of transaction costs is maybe the most noteworthy, as it implies that adjustment will occur only once deviations are higher than the transactions costs, and hence adjustment should not happen instantaneously and at each time. Financial theory predicts that even in highly liquid markets a so-called band of no arbitrage may exist where deviations are too small for the arbitrage to be profitable.

[11]See Stigler (2010) for an updated survey on threshold cointegration.

(Leybourne and Newbold, 2003) or in the variance of the innovation errors of each time series (Noh and Kim, 2003).

A nonlinear VECM may be denoted as

$$\Delta y_t = \begin{cases} A_1' X_{t-1}(\beta) + u_t & \text{if} \quad w_{t-1}(\beta) \leq \gamma \\ A_2' X_{t-1}(\beta) + u_t & \text{if} \quad w_{t-1}(\beta) > \gamma \end{cases}$$

with

$$X_{t-1} = \left\{ \begin{array}{ccccc} 1 & w_{t-1}(\beta) & \Delta x_{t-1} & \ldots & \Delta x_{t-n} \end{array} \right\}'$$

where x_t is a p-dimensional $I(1)$ time series which is cointegrated with one $(p \times 1)$ cointegrating vector β, $w_t(\beta) = \beta' x_t$ is the $I(0)$ error-correction term, u_t is an error term, A_1 and A_2 are coefficients matrices that describe the dynamics in each of the regimes, and γ is the threshold parameter.

The approach advocated by Balke and Fomby (1997) is to conduct a two-step analysis: If linear cointegration is not rejected, tests for threshold cointegration with linear under the null should be used. Failure of cointegration in the first step should lead to the use of tests with no cointegration under H_0 and threshold cointegration under the alternative. As the null hypothesis of none cointegrating relations is not rejected, we use the test developped by Seo (2006). The null hypothesis of no-linear cointegration is tested against the alternative of threshold cointegration. This test is superior to its precursors, such as those proposed by Balke and Fomby (1997) and Hansen and Seo (2002), where the standard test for linear cointegration is used for threshold cointegration.[12] The test statistic, denoted supW, is the supremum of the Wald test statistics for the null hypothesis, calculated from a grid of γ values over its parameter space. The statistic is denoted as

$$\text{supW} = \sup_{\gamma \in \Gamma} W_n(\gamma) \tag{8}$$

with γ the threshold parameter. Seo (2006) proves that, under certain conditions, the

[12] According to Seo (p.130, 2006), their approach is misleading and can cause substantial power loss.

supW converges to a function of the Brownian motion, free from nuisance parameters. For improved size and power properties in small samples, Seo (2006) proposes the bootstrap based on residual resampling, which approximates the sampling distribution of the supW statistic under the null. According to his Monte Carlo experiment, the bootstrap test shows desirable size properties and high power, especially when the sample size is as large as or more than 250. The result reported on Table 4 indicates that the null hypothesis of no cointegration cannot be rejected.

Thus, we can conclude that the spot and futures log-returns do not seem to be cointegrated.[13] As the two series are $I(1)$, we can estimate equation (4) to test the speculative efficiency hypothesis. As heteroskedasticity was found in the spot and futures prices (see Table 2), the models are estimated with the White heteroskedastic-consistent standard errors. Estimates of equation (4) is presented in Table 5. The results indicate that there is no evidence of a time-varying risk premiums ($\gamma = 0$). The null hypothesis $\delta = 1$ is rejected at the 5% level, implying that the futures prices of maturity 2010 and 2011 do not seem to be the best forecasts of the future spot prices. Nevertheless, the joint hypothesis of market efficiency and unbiasedness is not rejected, meaning that futures prices of maturity 2010 and 2011 appear to be the unbiased predictors of spot prices. [14] The result of the joint test is not the same of the futures series of maturity 2012. Indeed, the Wald test is rejected implying that futures prices of maturity 2012 are not the best predictors of spot prices. Our findings suggest that cointegration exists on the short term, not on the long-run. This phenomenon can be a particularity of the carbon markets which operates by phases. Market participants may consider 2012 as the end of the second phase and prefer waiting before taking any decision about 2012.

Nevertheless, the acceptance of the restrictions on parameters (γ and δ) are not a necessary and sufficient condition for market efficiency. The serial independence of

[13]Masih and Masih (2002) suggested that cointegration of commodity markets does not exist if there is either a non-stationary risk premium or a non-stationary convenience yield.

[14]The joint test is more powerful than the individual tests.

residuals is an important condition for market efficiency. Indeed, residual correlation implies that spot prices rely on past spot prices in addition to current futures prices, thus violating market efficiency. If the parameter restrictions and serial independence of residuals are met, then the market is efficient and futures prices provide unbiased estimates of future spot prices. The Breusch-Godfrey test is applied to check the presence of serial correlation (Table 5) and indicates that the null hypothesis of no serial correlation is rejected, which casts some doubt on the efficiency of the market. As the two above conditions are not met, the market does not seem to be efficient.

4 Conclusion

This paper investigated the speculative efficiency hypothesis on CO_2 emission allowance prices negotiated on Bluenext, by testing the joint hypothesis of market efficiency and unbiasedness of futures prices. The unit root tests concluded that spot and futures prices are non-stationary in levels but stationary in first-difference. The spot and futures prices are tested for cointegration using both linear (Johansen, 1995) and nonlinear (Seo, 2006) tests following the approach proposed by Balke and Fomby (1997). The results indicate the absence of linear and nonlinear cointegration relationship between spot and futures prices. The speculative efficiency hypothesis did not hold even if the restrictions on the parameters are not rejected because of the existence of serial correlation of residuals.

Understanding the relationship between spot and futures prices is thus of crucial importance for all participants in the carbon market. Carbon trading works only if markets for carbon provide enough liquidity and pricing accuracy, i.e., markets provide prices that are useful for hedgers and other users of carbon markets. The efficiency of the CO_2 market is particularly important for emission intensive firms, policy makers, risk managers and for investors in the emerging class of energy and carbon hedge funds. If carbon markets are inefficient the policy implications are that there is a greater role for regulation to improve information flows and reduce market manipulation (Stout, 1995). It is imperative that policy makers address these issues during the eminent reviewing process, to ensure that the EU ETS evolves into a mature, efficient and internationally competitive market.

Table 1: Results of unit root tests

Data	ERS	NP1	NP2	KPSS	KSS
Series in level					
Spot data	-0.220	-0.252	-0.217	1.431	-0.248
Dec10 Futures data	-0.067	-0.072	-0.068	1.592	-0.294
Dec11 Futures data	-0.152	-0.171	-0.148	1.633	-0.419
Dec12 Futures data	-0.158	-0.183	-0.153	1.667	-0.366
Series in first-difference					
Spot data	-2.368	-10.044	-2.233	0.319	-3.194
Dec10 Futures data	-2.791	-10.141	-2.739	0.298	-4.298
Dec11 Futures data	-3.425	-10.164	-2.902	0.276	-3.667
Dec12 Futures data	-3.355	-10.172	-2.926	0.249	-3.683
Critical value	-1.94	-8.10	-1.98	0.463	-2.93

Notes: The unit root tests are the efficient tests of Elliott, Rothernberg and Stock (1996, ERS) and Ng and Perron (2001, NP1 and NP2), the stationarity test of Kwiatkowski et al. (1992, KPSS), the nonlinear test of Kapetanios et al. (2003, KSS). ^{a} means significant at the 5% level.

Table 2: Statistical analysis of log-returns series

Data	Obs.	Mean (%)	SD	Skewness	Kurtosis	ARCH(10)
Spot data	623	-0.076	0.026	-1.194**	4.733**	15.303**
Dec10 Futures data	623	-0.084	0.025	-0.130	4.735**	16.680**
Dec11 Futures data	623	-0.086	0.025	-0.131	4.836**	15.652**
Dec12 Futures data	623	-0.087	0.024	-0.143	4.939**	14.405**

Notes: The skewness and kurtosis statistics are standard-normally distributed under the null of normality distributed returns. ARCH(10) indicates the Lagrange multiplier test for conditional heteroscedasticity with 10 lags. ** means significant at the 5% level. The futures data are the futures of maturity December 2010.

Table 3: Test for deterministic components in VECM

	Test statistic		
Null hypothesis	Dec10	Dec11	Dec12
Model 1 in model 2	-6.10	-4.75	-4.04
Model 2 in model 3	1.07	1.10	1.13
Model 3 in model 4	-6.76	-8.16	-8.34
Model 4 in model 5	2.15	2.75	3.70

Notes: The test statistic is asymptotically distributed as χ^2 with $(p-r)$ degrees of freedom with p the lag length and r the number of cointegration relationships. ** means significant at the 5% level.

Table 4: Cointegration tests

Hypotheses	*Johansen test*					
	Dec10		Dec11		Dec12	
	t-stat	p-value	t-stat	p-value	t-stat	p-value
Lambda max test						
$r \leq 0$	4.21	0.59	4.36	0.57	4.36	0.57
$r \leq 1$	0.11	0.79	0.27	0.67	0.41	0.59
Trace test						
$r \leq 0$	4.32	0.67	4.63	0.62	4.77	0.60
$r \leq 1$	0.11	0.79	0.27	0.67	0.41	0.59
	Seo test					
	Dec10		Dec11		Dec12	
	t-stat	p-value	t-stat	p-value	t-stat	p-value
no coint / threshold cointegration	20.71	1.00	18.09	1.00	18.93	1.00

Notes: ** means significant at the 5% level.

Figure 1: Spot and futures price series in log-returns

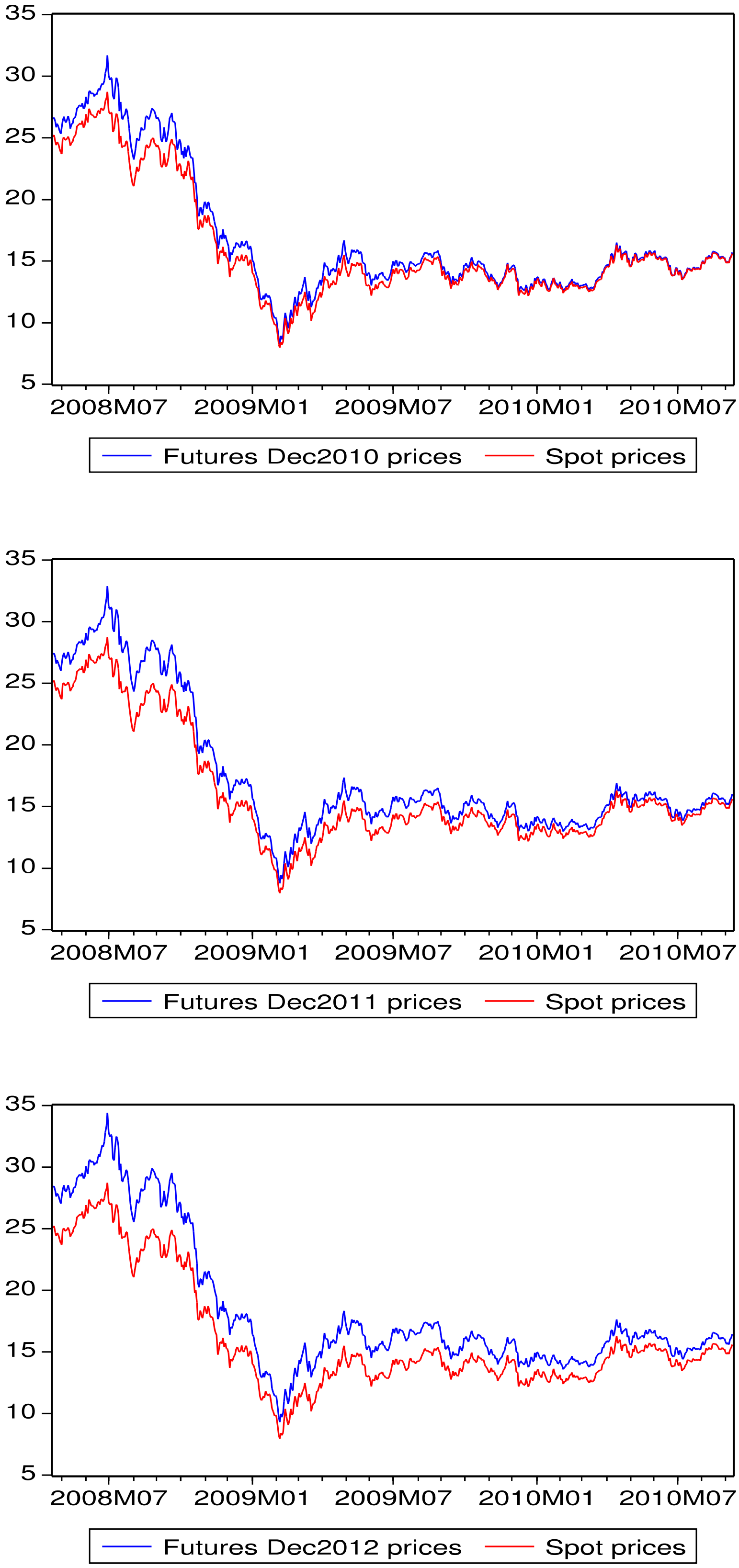

Table 5: Fama's regression

Model			Dec10	
	γ	δ	Wald	BG
$\Delta S_t = \gamma + \Delta F_t + \varepsilon_t$	7.0110^{-5} (0.80)	0.99^{**} (0.00)	0.75 (0.48)	22.20^{**} (0.00)
			Dec11	
	γ	δ	Wald	BG
$\Delta S_t = \gamma + \Delta F_t + \varepsilon_t$	1.0310^{-4} (0.72)	1.00^{**} (0.00)	0.07 (0.93)	14.41^{**} (0.00)
			Dec12	
	γ	δ	Wald	BG
$\Delta S_t = \gamma + \Delta F_t + \varepsilon_t$	-6.8810^{-4} (0.50)	0.09^{**} (0.03)	227.72 (0.00)	2.93^{**} (0.00)

Notes: ** means significant at the 5% level. BG represents the Breusch-Godfrey LM test for serial correlation. The test statistic is asymptotically distributed as χ^2 with p degrees of freedom where p represents the lag length.

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